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## Gender and corruption

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### Abstract

Using several independent data sets, we investigate the relationship between gender and corruption. We show using micro-data that women are less involved in bribery, and are less likely to condone bribe-taking. Cross-country data show that corruption is less severe where women hold a larger share of parliamentary seats and senior positions in the government bureaucracy, and comprise a larger share of the labor force. © 2001 Elsevier Science B.V. All rights reserved.

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### 1. Introduction

In recent years, there has been a concerted effort, by various national governments and international organizations, to increase the representation of women in public life. A prominent example is a recent legislation in France requiring all

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parties to field equal numbers of male and female candidates in all party-list elections, and equal numbers within an error margin of 2% in constituency-based elections (The Economist, 2000). Proponents of these reforms suggest that women may make different policy choices than men, and indeed, there is some evidence supporting this proposition.<sup>1</sup> Recently, however, an even more provocative claim has been made: in several different locations, influential public officials have advocated increasing representation of women on the grounds that this will lower the extent of corruption. In Mexico city, the police chief has taken away ticket-writing authority from the city's 900 male traffic policemen and created a new force consisting exclusively of women, hoping to reduce corruption (Moore, 1999). A similar policy has also been introduced in Lima, Peru where it is claimed there has been a fall in corruption after the introduction of women (McDermott, 1999). This paper evaluates the plausibility of such claims, using a variety of independent data sources.<sup>2</sup>

We present evidence that (a) in hypothetical situations, women are less likely to condone corruption, (b) women managers are less involved in bribery, and (c) countries which have greater representation of women in government or in market work have lower levels of corruption. This evidence, taken together, provides some support for the idea that, at least in the short or medium term, increased presence of women in public life will reduce levels of corruption.

Claims about gender differences can easily be misinterpreted. It is therefore important for us to clarify that we do *not* claim to have discovered some essential, permanent, or biologically determined differences between men and women. Indeed, the gender differences we observe may be attributable to socialization, or to differences in access to networks of corruption, or in knowledge of how to engage in corrupt practices, or to other factors. We do not attempt to identify these underlying factors, but rather to document several statistically robust relationships that point towards a gender differential in the incidence of corruption. We discuss

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<sup>1</sup> For example, Fukuyama (1998, p. 24) reports that the percentages of American women who supported US involvement in World War II, the Korean War, the Vietnam War, and the Gulf War, were 7 to 10 points less than the corresponding percentages for men. A study by the Center for American Women in Politics (Dodson and Carroll, 1991) documents substantial differences between men and women in their attitudes towards prohibition of abortion (79% women oppose vs. 61% men), towards the death penalty (49% women oppose vs. 33% men), and towards more nuclear plants (84% women oppose, compared to 71% men). Hun and Jones (1999) find that there are significant differences by gender in committee membership and bills introduced to the Argentinean legislature, with women over-represented on health and education committees and under-represented on finance, defense, and foreign policy.

<sup>2</sup> Kaufmann (1998) presents a scatterplot showing a cross-country correlation between corruption and an index of women's rights and emphasizes the need for more detailed investigation of this association.

some theories about the origin of this differential in Section 5 below, but do not take a position on this question.

Our evidence is organized as follows. We first present data from the World Values Survey, in which men and women in a large number of developed and developing countries were asked a series of questions regarding their attitudes in hypothetical situations in which there was room for dishonest or opportunistic behavior. We show that men were more likely to choose options that are equivalent to the “defect” option in a prisoners’ dilemma game. After showing gender differences in a range of attitudes, we present more detailed multivariate evidence on gender differentials in the attitude to bribery. We then present evidence of behavior in actual as opposed to hypothetical situations. Using a survey of enterprise owners and managers in Georgia (formerly part of the Soviet Union), we show that officials in firms owned or managed by men are significantly more likely to be involved in bribe-giving.<sup>3</sup>

One concern in the above analyses is that corruption is self-reported. Because of this, it is conceivable that our results reflect gender differentials in *acknowledgment* of corruption, rather than in *incidence* of corruption. Data on corruption which are not self-reported are available only at the national level. Using corruption indices developed by Kaufmann et al. (1999), Transparency International, and Political Risk Services, we find that greater participation by women in market work and government is associated with lower levels of corruption. This result is of value not only because national-level corruption data are not self-reported, but also because it shows that gender differentials have macro-level impacts. These findings are consistent with arguments that at least in the short run policies designed to increase the role of women in commerce and politics, commonly justified on grounds of gender equity and poverty alleviation, may also have an efficiency payoff, by lowering corruption.<sup>4</sup>

## 2. Micro-evidence: the World Values Surveys

The World Values Surveys are a set of surveys carried out in dozens of developed and developing countries in the early 1980s and the early 1990s. The purpose of these surveys was to collect information on the attitudes and values of the peoples of various societies around the world. An effort was made to ensure

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<sup>3</sup> We chose Georgia (rather than any other country) purely because of access to micro-data on corruption: the World Bank was kind enough to allow us to use this survey.

<sup>4</sup> The adverse consequences of corruption have been discussed by Klitgaard (1988), Knack and Keefer (1995), Mauro (1995, 1998), and Olson et al. (2000).

that in each case, the sample was nationally representative.<sup>5</sup> We are able to use data from 18 surveys in 1981 and 43 surveys in 1990–1991.

In addition to hundreds of other items, these surveys inquire about the acceptability of various dishonest or illegal behaviors. For each behavior, respondents are asked to place themselves on a 1–10 scale, where 1 indicates that the behavior can “never be justified” and 10 indicates it can “always be justified”. For most items in most countries, the natural cut-off point is at the value 1, as a majority of respondents typically assert (fortunately) that the behavior can never be justified. Aggregating over all countries in the surveys, the gender gap consistently favors women, as shown in Table 1a. For all 12 items listed, a significantly higher percentage of women than men believe that the illegal or dishonest behavior is never justifiable. The gap ranges from more than nine percentage points for driving under the influence to about four points for claiming government benefits for which one is ineligible. In all cases, the gender differences are significant at the 0.0001 level.

The case of greatest interest to us is “someone accepting a bribe in the course of their duties”: 77.3% of women but only 72.4% of men agree that this behavior is “never justified”. This difference implies that about one-fifth more men than women (27.6% compared to 22.7%) believe that bribery can sometimes or always be justified. Table 1b provides a breakdown of percentages across the 1–10 scale. As mentioned above, females are more concentrated at category 1. The pattern is then slightly reversed for all the categories 2 through 10, which is not surprising, since each set of percentages must total 100.

This comparison of proportions could be misleading if men and women differed systematically in some other characteristic that also affects the attitude to bribery. In Table 2a, b, and c, we show that this result is robust to tests that control for other respondent characteristics. In the first column of Table 2a, our dependent variable takes the value 1 if the respondent says that bribery is “never justified”, and zero otherwise. Our main interest is in the coefficient on the gender dummy (1 if male). There is some evidence that rule-breaking is higher among young people, so we include age as a regressor. Marriage is often believed to alter public behavior; this is reflected, for instance, in lower rates of incarceration among married men, as compared to single men (Akerlof, 1998). To account for this, we include a dummy, which takes the value 1 if the respondent is married. Commit-

<sup>5</sup> Inglehart et al. (1998) provide details on the procedures followed in the various surveys in the 1990s. The surveys in the western countries were carried out by experienced survey organizations, many linked with the Gallup chain. In other countries, they were carried out by academies of science, or by university-based institutes. Inglehart et al. (p. 471) write: “In most countries stratified multistage random sampling was used, with samples selected in two stages. First a random selection of sample locations was made ensuring that all types of location were represented in proportion to their population. Next, a random selection of individuals was drawn up.”

Table 1

(a) Gender and socially cooperative attitudes, World Values Surveys

	% Saying the behavior "can never be justified"	
	Male	Female
(1) Claiming government benefits which you are not entitled to	63.7	67.9
(2) Avoiding a fare on public transport	60.3	64.9
(3) Cheating on taxes if you have the chance	54.4	61.5
(4) Buying something you knew was stolen	72.9	79.5
(5) Taking and driving away a car belonging to someone else	83.1	87.2
(6) Keeping money that you have found	43.9	51.6
(7) Lying in your own interest	45.1	50.9
(8) Someone accepting a bribe in the course of their duties	72.4	77.3
(9) Fighting with the police	52.0	57.1
(10) Failing to report damage you've done accidentally to a parked vehicle	61.8	67.6
(11) Throwing away litter in a public place	69.1	74.4
(12) Driving under the influence of alcohol	74.2	83.4

Sample sizes vary between 52,107 and 83,532. All differences are significant at the 0.0001 level.

(b) Gender differentials in the attitude towards bribery

	Female (%)	Male (%)
1 – Bribery is never justified	77.33	72.39
2	8.44	9.17
3	5.05	6.07
4	2.44	3.03
5	2.81	4.03
6	1.36	1.83
7	0.76	1.06
8	0.56	0.82
9	0.35	0.45
10 – Bribery can always be justified	0.90	1.15

ment to a religion is often believed to affect behavior;<sup>6</sup> therefore, we include a dummy which takes the value 1 if the individual responded yes to the question "are you a religious person?" We also include another dummy which takes the value 1 if the respondent frequently attends religious services. Finally, the

<sup>6</sup> This could be because, as argued by Strate et al. (1989), "...Church attendance involves a sense of personal affiliation with an institution in which communal values and social obligations are regularly emphasized."

Table 2

(a) Gender differentials in the attitude towards bribery: probit and ordered probit estimates (S.E. in parentheses)

	Probit (dependent variable = 1 if the response was 1 meaning bribery is "never justified"; = 0 otherwise)	Ordered probit (dependent variable is an integer between 1 and 10, where 1 corresponds to bribery is "always justified"; 10 = "never justified")
Male	-0.140 (0.010) **	-0.142 (0.010) **
Marginal effect (%)	-4.3	-4.0
School 16	0.039 (0.013) **	0.062 (0.012) **
Married	0.067 (0.012) **	0.072 (0.011) **
Attend religious services often	0.096 (0.014) **	0.095 (0.014) **
Religious	0.135 (0.012) **	0.122 (0.011) **
Age	0.025 (0.002) **	0.023 (0.002) **
Age squared	-0.00013 (0.00002) **	-0.00012 (0.00002) **
Constant	-0.507 (0.056) **	
No. of observations	79,645	79,645

(b) Country-wise probit regression, World Values Survey data, 1990–1991 [Dependent variable = 1 if bribery is "never justified", = 0 otherwise (S.E. in parentheses)]

Country	No. of observations	Male dummy coefficient	S.E.	Marginal effect (%)
Sweden	963	-0.360	(0.090) **	-11.6
Estonia	999	-0.327	(0.087) **	-11.5
Slovenia	1003	-0.427	(0.096) **	-11.4
Latvia	890	-0.367	(0.099) **	-11.4
Netherlands	1002	-0.304	(0.087) **	-10.3
Bulgaria	1007	-0.336	(0.093) **	-9.2
Austria	1445	-0.251	(0.073) **	-8.5
Mexico	1497	-0.200	(0.067) **	-7.9
Japan	969	-0.227	(0.088) **	-7.8
Switzerland	1356	-0.291	(0.082) **	-7.4
France	983	-0.197	(0.085) **	-7.3
S. Africa	2675	-0.253	(0.057) **	-7.1
W. Germany	2049	-0.180	(0.059) **	-6.8
Canada	1695	-0.215	(0.069) **	-6.5
Russia	1909	-0.279	(0.075) **	-6.3
Iceland	679	-0.246	(0.122) *	-5.7
Denmark	995	-0.367	(0.119) **	-5.6
E. Germany	1329	-0.140	(0.073) *	-5.2
Belgium	2671	-0.131	(0.050) **	-5.1
Ireland	996	-0.220	(0.103) *	-4.8
Britain	1478	-0.132	(0.074) *	-4.0
N. Ireland	303	-0.182	(0.189)	-3.9

Table 2 (Continued)

(b) Country-wise probit regression, World Values Survey data, 1990–1991 [Dependent variable = 1 if bribery is “never justified”, = 0 otherwise (S.E. in parentheses)]

Country	No. of observations	Male dummy coefficient	S.E.	Marginal effect (%)
Norway	1226	0.134	(0.085)	–3.6
Finland	580	–0.104	(0.115)	–3.3
Brazil	1761	0.169	(0.080)*	–3.3
Hungary	979	–0.085	(0.085)	–3.2
Moscow	980	–0.094	(0.090)	–3.1
China	988	–0.151	(0.106)	–3.0
Argentina	986	–0.285	(0.139)*	–2.9
Spain	4048	–0.093	(0.046)*	–2.6
Turkey	1013	–0.153	(0.141)	–2.2
Portugal	1148	–0.061	(0.083)	–2.0
Czech–Slovak	1392	–0.049	(0.069)	–1.9
Chile	1477	–0.079	(0.082)	–1.9
USA	1754	–0.032	(0.071)	–0.9
Poland	924	–0.031	(0.101)	–0.8
Italy	2014	0.008	(0.063)	0.2
Nigeria	979	0.024	(0.090)	0.9
India	2476	0.052	(0.064)	1.2
S. Korea	1245	0.082	(0.088)	1.9
Lithuania	978	0.055	(0.086)	2.1
Belarus	997	0.192	(0.089)*	6.4
Romania	1088	0.220	(0.083)* *	7.9

(c) Country-wise probit regression, World Values Survey data, 1981 [Dependent variable = 1 if bribery is “never justified”, = 0 otherwise (S.E. in parentheses)]

Country	No. of observations	Male dummy coefficient	S.E.	Marginal effect (%)
Netherlands	1153	–0.355	(0.079)* *	–13.1
Belgium	1023	–0.279	(0.082)* *	–10.6
Japan	1077	–0.272	(0.081)* *	–9.7
France	1161	–0.213	(0.078)* *	–8.4
Iceland	904	–0.336	(0.106)* *	–8.1
Sweden	938	–0.252	(0.091)* *	–7.9
Argentina	879	–0.197	(0.102)*	–5.1
Denmark	1175	–0.370	(0.117)* *	–4.6
Spain	2205	–0.125	(0.063)*	–3.7
Norway	1228	–0.141	(0.088)	–3.5
N. Ireland	310	–0.207	(0.204)	–3.2
Britain	1182	–0.103	(0.085)	–3.0
Italy	1305	–0.077	(0.076)	–2.8
W. Germany	1301	–0.061	(0.073)	–2.3
USA	2270	–0.071	(0.062)	–1.9

(continued on next page)



Table 2 (continued)

(c) Country-wise probit regression, World Values Survey data, 1981 [Dependent variable = 1 if bribery is "never justified", = 0 otherwise (S.E. in parentheses)]

Country	No. of observations	Male dummy coefficient	S.E.	Marginal effects
Australia	1176	−0.047	(0.086)	1.4
Canada	1240	−0.042	(0.082)	1.3
Ireland	1145	−0.014	(0.088)	0.4

<sup>a</sup>The scale for this dependent variable was reversed to produce estimates comparable to the others.

<sup>\*</sup> Significant at 5% for two-tailed tests.

<sup>\*\*</sup> Significant at 1% for two-tailed tests.

education dummy takes the value 1 if the respondent was schooled beyond age 16.<sup>7</sup>

In the first column of Table 2a, we pool the data across countries and estimate a probit model. In order to control for unobservable country characteristics that might otherwise bias our results, we include a dummy for each survey.<sup>8</sup> The coefficient on the gender dummy (1 if male) is negative and is statistically significant at any reasonable level. The marginal effect corresponding to this coefficient is −4.3%, i.e. all else being equal, a man's likelihood of responding that accepting a bribe is "never justified" is 4.3 percentage points less than the likelihood for a woman.<sup>9</sup> Though this differential is not very large, it occurs with remarkable consistency, as can be seen below.

In column 2, we estimate an ordered probit model, using categories 1 through 10. The scale was reversed to allow comparability with the probit model reported above. The male dummy coefficient is again negative and statistically significant. The marginal effect corresponding to this coefficient is −4%, i.e. all else being equal, a man's likelihood of responding that accepting a bribe is "never justified" is four percentage points less than the likelihood for a woman.

It is possible that a large gender differential in a subset of countries is driving the results in Table 2a. Therefore, we estimated the probit model separately for each country and found, in Table 2b and c, that the gender differential is observed

<sup>7</sup> Inclusion of additional education dummies did not alter our central result regarding the gender differential. The data do not allow us to construct a variable equal to years of education completed.

<sup>8</sup> So, for instance, there are separate dummies for Canada in 1981 and 1991.

<sup>9</sup> We re-estimated the probit model perturbing the dependent variable in two ways: (a) dependent variable = 1 if the respondent chose categories 1 or 2 and (b) dependent variable = 1 if respondent chose categories 1, 2, or 3. We saw in the summary (Table 1b) that while a larger percentage of women are in category 1, the gender differential is slightly reversed for categories 2 through 10. Consistent with this, we found that while the male dummy is negative and significant, the marginal effects are slightly smaller (less negative), at −3.5% and −2.7% for cases (a) and (b), respectively.

in most countries, although the estimated effects vary. We estimated probit models for 43 countries for 1991 and 18 countries for 1981, using the same specification used in column 1 of Table 2a. In 1991, we see that in 36 of 43 countries, the gender differential favors women; in 24 of these countries, the differential is statistically significant at 5%. There are only seven countries in which the gender differential favors men and only two of these differentials are statistically significant at 5%. In the data from 1981 (Table 2c), the gender differential favors women in all 18 countries; the differential is statistically significant at 5% in nine of these. Thus, the gender differential in the attitude to corruption seems to be a more or less a worldwide phenomenon.

Some readers have pointed out that the male dummy is negative and significant in only slightly more than half the cases. This is a fair point. On the other hand, it should be noted that if there were no gender differentials in any country, the probability of getting 54 or more negative signs out of 61 is virtually zero.

A possible response to these findings is that women may disapprove of corruption more only because they are less likely than men to be employed.<sup>10</sup> Persons not employed may be less able to benefit from corruption, or norms regarding bribery may be different among employed and non-employed persons (e.g. the latter may be more naïve or idealistic). Accordingly, we re-estimated the model in column 1 of Table 2a, including a dummy variable for the employment status of the individual (1 if employed). The coefficient on the male dummy became larger (more negative), indicating that the gender differential in attitudes is not an artifact of male–female differences in employment rates.

The above analysis is based on attitudes towards the acceptability of taking bribes. In the following section, we present evidence of actual behavioral differences in bribe-paying, with respect to gender, drawing on an enterprise survey in Georgia.

### 3. Micro-evidence: an enterprise survey in Georgia

In this section, we use data from a World Bank study of corruption in Georgia, which included a survey of 350 firms.<sup>11</sup> The firms were in four broad sectors: trade, manufacturing, services and agriculture. We categorize them in three groups: large (more than 50 employees), medium (between 10 and 50 employees), and small (less than 10 employees). The incidence of corruption is high, as firms reported paying an average of 233 lari per month (1.5 lari is equivalent to 1 US dollar) or 9% of turnover on average (Anderson et al., 1999).

<sup>10</sup> We thank Margaret Madajewicz for this point.

<sup>11</sup> We are grateful to the World Bank for making these data available to us.

Table 3

## (a) Means, Georgia survey

	Unit	Whole sample <sup>a</sup> ( <i>n</i> = 2219)	Male owner/ senior manager ( <i>n</i> = 1717)	Female owner/ senior manager ( <i>n</i> = 502)
Frequency of bribes	Percent	10.7% (S.D. = 26.9)	12.5% (S.D. = 28.9)	4.6% (S.D. = 17.5)
<i>Size of firms</i>				
Small	Dummy	0.48	0.42	0.67
Medium	Dummy	0.33	0.35	0.27
<i>Ownership</i>				
Majority state ownership	Dummy	0.1	0.12	0.03
Foreign participation	Dummy	0.38	0.44	0.19
<i>Sector</i>				
Trade	Dummy	0.55	0.51	0.68
Manufacturing	Dummy	0.25	0.30	0.08
Services	Dummy	0.44	0.44	0.43
<i>Education of senior manager</i>				
University	Dummy	0.82	0.83	0.80
Post-university	Dummy	0.07	0.08	0.01
<i>Scope of operation</i>				
Local	Dummy	0.64	0.59	0.82
Regional	Dummy	0.06	0.06	0.07
National	Dummy	0.15	0.17	0.09
Percentage of sales domestic	Percent	95.41	94.18	99.6

## (b) Georgian enterprises, patterns of bribe-paying

Type of procedure	Probit	Probit marginal effect (%)
Male owner/ senior manager	0.674 * * (0.190)	12.9
<i>Size of firms</i>		
Small (reference group = large)	1.621 * * (0.326)	39.0
Medium	1.031 * * (0.293)	28.1
<i>Ownership</i>		
Majority state ownership	−0.791 * (0.326)	−12.9
Foreign participation	−0.006 (0.190)	−0.2
<i>Sector</i>		
Trade (reference group = agriculture)	0.009 (0.174)	0.2
Manufacturing	0.194 (0.203)	4.8
Services	0.253 (0.172)	6.0

Table 3 (continued)

(b) Georgian enterprises, patterns of bribe-paying

Type of procedure	Probit	Probit marginal effect (%)
<i>Education of senior manager</i>		
University (reference group - below university)	-0.264 (0.238)	-6.7
Post-university	0.061 (0.396)	1.5
<i>Scope of operation</i>		
Local (reference group - international)	-1.175 * (0.488)	-31.6
Regional	-1.592 * * (0.515)	-16.9
National	-1.148 * (0.500)	-17.3
Percentage of domestic sales	0.012 (0.009)	0.28
Constant	-2.561 * * (0.724)	
No. of observations	2219	
Pseudo $R^2$	0.165	

(1) Dummies were included for the agency with which firm was in contact.

(2) Standard errors have been corrected for within-firm autocorrelation of error terms.

(3) Marginal effects were computed at the sample means.

<sup>a</sup>The sample of firms in which the owner/senior manager was interviewed. The proportions in various sectors add up to more than 100% because some firms are in more than one sector (say, Trade and Manufacturing).

\* Significant at 5%.

\*\* Significant at 1%.

Managers were asked about contact with and illegal payments to 18 different agencies.<sup>12</sup> There are potentially 6300 (350 firms  $\times$  18 agencies) points of contact, but only 2954 of these actually occurred. To maximize the reliability we used only data on firms where the senior manager/owner was interviewed, which left 2219 observations.<sup>13</sup> Summary statistics, by gender, are provided in Table 3a.

Our analysis starts with the response to the following question: "How frequently do the officials providing the service require unofficial payments? Please answer on a scale of 1 to 7, where 1 = never, 2 = 1–20% of the time, 3 = 21–40%

<sup>12</sup> The full list of contact agencies is as follows: phone installation, enterprise registration, water, electricity, inspection of weights and measurements, fire inspection, sanitary inspection, tax and finance inspection, tax clearance (for example, in government privatization), other clearance to participate in government procurement, export license or permit, import license or permit, customs at border crossing, registration of property ownership, lease of state-owned commercial real estate, state banking services, building permits and road police.

<sup>13</sup> Results are very similar if we use the full sample; see below.

of the time, 4 = 41–60% of the time, 5 = 61–80% of the time, 6 = 81–99% of the time, and 7 = always.” Firms owned or managed by women gave bribes on average on 4.6% of the occasions that they came in contact with a government agency; the average percentage was more than twice as large for firms owned/managed by men, 12.5%.<sup>14</sup> Thus, the descriptive evidence is strongly suggestive of a gender differential in involvement in bribery.<sup>15</sup>

How should this evidence be interpreted? The way the question is phrased, it appears that the impetus for the bribe is coming from the official, not from the owner/manager. However, questions on bribery are usually put in this way to avoid placing the onus of the bribe on the respondent, in the hope of eliciting an honest response. Therefore, an obvious interpretation of these results is that female owners/managers are less likely to offer bribes than male owners/managers. However, other interpretations are possible. It could be that women are less likely to belong to bribe-sharing old boy networks, and hence, may be less prone to be asked for bribes. It could also be that, due to less individual or collective experience in the labor force, women have not yet “learned” how to engage in corruption. Here, we document the presence of a statistically robust gender differential, but do not attempt to distinguish among these alternative interpretations.

Table 3b examines whether this gender differential remains after we control for other firm characteristics. Given there are seven categorical outcomes which can be meaningfully ranked, one possibility is to estimate an ordered probit model. The dependent variable here takes values 1 through 7, with 1 being the category “never”. However, if we are only interested in the distinction between firms, which never give bribes and those which sometimes do, a probit model is appropriate. Here, the dependent variable takes the value 0 if the firm never gives a bribe, and 1 if it sometimes does.

We have relied on the literature on Georgia, and on corruption more broadly, to guide our choice of control variables. Since a firm’s size can affect its ability to pay, as well as its bargaining power or “connections”, we include size dummies (small and medium, with large being the excluded category). For similar reasons, we include dummies to reflect the firm’s scale of operations (local, regional, and national, with international being the excluded category), and the percentage of the firm’s output sold domestically.

Depending on the sector in which the firm operates its dependence on governmental services and hence, its temptation to bribe, may vary; therefore, we have included sector dummies (manufacturing, services, and trade, with agriculture

<sup>14</sup> A value in the range 21–40% was converted to 30%, and so on. We then took the average of these converted values by gender.

<sup>15</sup> If we use the full sample, the average percentages for firms run by men and women are 10% and 4.1%, respectively.

being the excluded category). We also include dummies for the level of education of the owner/manager; these could partially reflect influence or connections as, for example, in “old boy networks”. The dummies are for university and post-university, with the excluded category being those who do not have a university education. Since some governmental agencies are likely to be more corrupt than others, we include dummies for the agency with which the firm is having contact. Because these dummies are so numerous (18 agencies, hence, 17 dummies) these coefficients are not reported.<sup>16</sup> Participation by the state and foreign participation could also affect bribe-giving, and dummies are included for these.<sup>17</sup>

Column 1 of Table 3b presents probit estimates; as mentioned above, the dependent variable takes the value 0 if the firm is in the “never” category and 1 otherwise. The male dummy has a positive (and statistically significant) coefficient which suggests that, all else being equal, a firm owned/managed by a man is less likely to be in the “never” category than a firm owned/managed by a woman. Column 2 presents marginal effects, i.e. the effect of a unit increase in the explanatory variable on the probability that the firm is not in the “never” category. We see that the marginal effect of a male owner/manager is 12.9% percentage points.<sup>18</sup> We also see that firm size, state ownership, and the scope of operations of the firm have large (and statistically significant) impacts on the probability that the firm is in the “never” category. When we estimate an ordered probit model, the marginal effect of a male owner/manager (defined in the same fashion as above) is 13.7%.

#### 4. Macro-evidence: cross-country tests

Having seen evidence from micro-data based surveys, in the next section we turn our attention to analysis of country-level data. These cross-country analyses complement the micro-level evidence in two important ways. First, as earlier mentioned, national-level corruption ratings are not self-reported, so that any gender differentials cannot be produced by male–female differences in the willingness to acknowledge corruption. Second, micro-level evidence carries no necessary implications for the macro-level relationship between women’s participation and the severity of corruption in public life. For example, male–female differences

<sup>16</sup> The most corrupt agencies in terms of frequency of receiving bribes are traffic police, customs, import/export licensing and tax inspection. The amounts of the bribes are relatively large for customs, import/export licensing, building permits, tax inspection and enterprise registration.

<sup>17</sup> State-ownership should reduce bribe-giving if this gives the firm better contacts within the government. Miller et al. (1999) report that in formerly communist countries, officials treat other officials better than they do private citizens. Foreign ownership may increase bribe-giving, since foreign-owned firms may be perceived to be richer, and more able to pay.

<sup>18</sup> Estimating a probit model using the full sample yields a slightly smaller gender differential, 9.7%.

in attitudes and behavior may be too small for an increase in women's participation in commerce and government to move society from a highly corrupt to a less corrupt equilibrium or women may have little influence on the way public life is conducted, even when their participation rates rise, as long as they remain in the minority.

#### *4.1. Measurement and data*

No objective measure of corruption with broad cross-country coverage is available, so we rely on subjective indicators. The most widely known measure of corruption is Transparency International's "Corruption Perceptions Index". This index combines the information available from numerous sources, some using investor surveys and others based on assessments of country experts. The TI index can vary between 0 and 10, with higher values signifying less corruption.<sup>19</sup>

Kaufmann et al. (1999) construct a similar index, using data largely from the same sources. Their "graft index" differs from TI's index in two major ways. First, rather than weighting all available sources equally, their statistical procedure assigns lower weights to sources that tend to agree less closely with other sources. This difference in the way the graft index and the TI index are constructed has little impact, as the two indexes are correlated at 0.98. A second and more important difference between them is that the graft index covers more countries. It is therefore used as the primary corruption measure in this section, but the main tests using a 93-country sample for the graft index are replicated using a 68-country sample for the TI index. The graft index is constructed to have a mean value of 0 and a standard deviation of 1 in the full Kaufmann et al. (1999) sample.

The graft index and the TI index take into account both "grand" or high-level corruption as well as petty corruption, as indicated by the criteria used in the corruption ratings provided by the International Country Risk Guide (ICRG), one of the various sources used in constructing both indexes. Lower scores by ICRG indicate that "high government officials are likely to demand special payments", and that "illegal payments are generally expected throughout lower levels of government" in the form of "bribes connected with import and export licenses, exchange controls, tax assessment, police protection, or loans". The fact that the graft index and TI index measure a combination of grand and petty corruption has implications for the ways in which women's participation should be measured in cross-country tests.

The firm-level analysis in the preceding section focused on the gender of the owner/manager. For country-level analyses, several relevant measures of women's involvement in politics and commerce are available: the proportion of legislators

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<sup>19</sup> Details of the method of construction are provided by Lambsdorff (2000).

Table 4  
Summary statistics for 93-country sample

Variable	Mean	Standard deviation	Minimum	Maximum
Graft index	0.17	0.95	– 1.57	+ 2.13
TI index ( $N = 68$ )	5.03	2.43	1.5	10
Women in parliament (%)	9.7	8.2	0	39
Women government ministers (%)	7.8	8.3	0	44
Women in labor force (%)	38.5	7.8	10.7	53.9
Women's influence index	– 0.01	0.79	– 1.91	3.02
Log (GNP per capita, 1995)	8.24	1.10	5.95	9.96
Average years of schooling, 1990	5.63	2.69	0.65	11.74
Catholic proportion	34.3	37.2	0	100
Muslim proportion	20.2	34.4	0	99.8
Former British colony	0.38	0.49	0	1
Never colonized	0.26	0.44	0	1
Largest ethnic group (%)	70.6	24.2	17	100
Political freedoms	4.87	1.97	1	7

in the national parliament who are female,<sup>20</sup> the proportion of ministers and high-level government bureaucrats who are women, and women's share of the labor force. Table 4 provides summary statistics for the variables used in the cross-country tests.

Our three measures of women's participation and a composite measure we constructed (described below) can, at one level, be considered proxies for the overall participation of women in politics and commerce. However, they can also be given more specific interpretations. The share of women in parliament can affect corruption levels in at least two ways. First, legislative corruption is itself an important dimension of governmental corruption, and if women tend to accept fewer bribes, the incidence of legislative corruption will be lower where women hold more seats. Second, members of parliament may influence the incidence of bureaucratic and judicial corruption through the passage of laws designed to deter bribery, through their influence on judicial or executive branch appointments (in some countries), or through placing corruption on the public agenda and encouraging the media and other elements of civil society to focus on the problem.

The share of ministers and top-ranking bureaucrats is a supplementary measure of women's participation in politics.<sup>21</sup> The incidence of bribe-taking in high-level

<sup>20</sup> This measure is based on legislators in both houses of parliament, for countries that have upper and lower chambers. All results using this measure are robust to using alternatively the proportion of lower-house members who are women.

<sup>21</sup> Included are cabinet ministers, deputy and vice ministers, permanent secretaries, deputy permanent secretaries and heads of Central Banks. The data are published in the UN's 1999 Human Development Report, and are collected from the *Worldwide Government Directory*, published by Worldwide Government Directories, Bethesda, MD.



positions in the bureaucracy may be reduced where more of those positions are held by women. “Petty” corruption at lower levels of the bureaucracy can also be affected, to the extent that ministers and sub-ministers select lower-level government officials and influence the formulation and enforcement of rules against bribe-taking.

Although women’s share of elite positions can influence petty corruption, it is useful to have a supplementary measure of women’s representation in lower levels of the government bureaucracy as well as in the private sector. Data on the share of lower-level government positions held by women are unavailable. Women’s share of the overall labor force is the closest available proxy. Women’s share of the labor force is likely to also capture, to some extent, any tendency for women in the private sector to offer bribes less frequently than men.

Not surprisingly, where women are better represented in parliament, they also tend to be better represented in top ministerial/bureaucratic positions, and even in the labor force more generally. Women’s share in parliament is correlated with women’s share of top ministerial/bureaucratic positions at 0.74 ( $p = 0.0001$ ), and with labor force share at 0.33 ( $p = 0.0015$ ). The latter two variables are correlated with each other at 0.26 ( $p = 0.011$ ). Therefore, when only one of these three variables at a time, for example, women in parliament, is included in a corruption regression, its coefficient captures at least part of the effects of other dimensions of women’s influence. Accordingly, we also report tests using an index of women’s influence, which incorporates all three variables.

Our tests of the relationship between the level of corruption and women’s participation control for many other potential determinants of corruption. We control for (the log of) per capita income for two reasons. First, the development of institutions to restrain corruption may be a costly activity undertaken more easily by richer countries. Second, in some cases where survey respondents have little concrete information on which to base their assessments, they may simply infer that corruption is a problem where they observe incomes to be low. To the extent that formulation, implementation, and public knowledge of written codes and laws reduce corruption, a more educated population may be less tolerant of corruption. Therefore, we control for the average years of education completed by adults, using data from Barro and Lee (1993). Percent of the population who are Catholic and percent of the population who are Muslim are included as proxies for “cultural” factors that may affect women’s participation and/or corruption. For example, casual observation suggests that within Europe, Catholic countries such as Italy and Spain have lower rates of women’s participation and more severe corruption than the Protestant Scandinavian nations.<sup>22</sup>

<sup>22</sup> Disputes at the 1995 UN women’s conference in Beijing and in a special UN session on gender and development in June 2000 usually pitted Western delegates against delegates from predominantly Catholic and Islamic nations. (See, for example, *The Washington Post*, 2000 [June 10, 2000, A20]).

Corruption may also be linked to the history of colonialism. Therefore, we include a dummy, which takes the value 1 if the country has never been a colony. It has also been argued that the character of British colonialism was different from others, so we include a dummy (1 if former British colony) to allow for this possibility.<sup>23</sup> Corruption may be higher in more ethnically divided societies. Therefore, we include the percentage of population belonging to the largest ethnic group as a regressor, using data from Sullivan (1991).

Democratic political institutions can restrain corruption in several ways. Multi-party competition may reduce corruption because each party has the incentive to expose any wrongdoing by another party.<sup>24</sup> By increasing the threat of exposure, an independent media can increase the costs of corrupt behavior. Independent judiciaries may reduce the incidence of corruption, at least within the executive branch. As a summary measure of democratic institutions that can restrain corruption, we use the well-known Freedom House political freedoms' indicator, using ratings for 1995.<sup>25</sup> Values range from 1 to 7; following common practice, we reverse the original scale so that higher values indicate more political freedoms.

## 4.2. Results

Estimates from cross-country tests using the graft index are presented in Table 5. In Eq. (1), women's share of parliamentary seats is highly significant; the coefficient implies that a one standard deviation increase (about eight percentage points) is associated with an increase in the graft index of slightly more than one-fifth of a standard deviation. By comparison, a standard deviation increase in (the log of) per capita income is associated with an increase in the graft index of slightly more than one half of a standard deviation. Other significant variables include the former British colony dummy and the political freedoms index. Other things equal, ex-British colonies score nearly half a standard deviation higher on

<sup>23</sup> Using the TI index, Treisman (1998) finds that ex-British colonies are rated as less corrupt on average.

<sup>24</sup> For a discussion of this and related issues, see Shleifer and Vishny (1998) and Myerson (1999).

<sup>25</sup> The Freedom House civil liberties index is equally relevant as the political freedoms index, but the former is not used because it is correlated by construction with corruption, as it includes "freedom from gross government indifference and corruption" among its evaluative criteria. The political freedoms and civil liberties indexes for 1995 are correlated at 0.89 in our sample. In tests not reported in tables, we alternatively controlled for corruption-restraining institutions using several indexes from Humana (1992) which evaluate the independence of the courts, the degree of multiparty competition, press censorship, independence of newspapers, and independence of TV and radio. The major difference with results using the political freedoms index is a reduction in the sample size; coefficients on the women's participation variables are unaffected.

Table 5  
Determinants of corruption, cross-country regressions (Dependent variable: graft index)

Equation					
	1	2	3	4	5
Parliament, proportion women	2.456 <sup>***</sup> (0.751)			1.273 (0.853)	
Ministers, proportion women		2.432 <sup>***</sup> (0.567)		1.444 (0.813)	
Labor force, proportion women			2.419 <sup>***</sup> (0.767)	2.048 <sup>*</sup> (0.804)	
Women's influence index					0.364 <sup>***</sup> (0.062)
Log (GNP per capita, 1995)	0.478 <sup>***</sup> (0.090)	0.459 <sup>***</sup> (0.090)	0.567 <sup>***</sup> (0.098)	0.551 <sup>***</sup> (0.098)	0.532 <sup>***</sup> (0.087)
Average years of schooling, 1990	0.003 (0.035)	0.009 (0.037)	0.012 (0.040)	0.003 (0.035)	0.003 (0.035)
Catholic proportion	-0.281 (0.159)	-0.354 <sup>*</sup> (0.172)	-0.221 (0.179)	-0.139 (0.163)	-0.167 (0.164)
Muslim proportion	-0.152 (0.192)	-0.192 (0.197)	-0.066 (0.196)	0.017 (0.198)	-0.014 (0.190)
Former British colony	0.481 <sup>***</sup> (0.131)	0.467 <sup>***</sup> (0.133)	0.418 <sup>***</sup> (0.134)	0.469 <sup>***</sup> (0.127)	0.476 <sup>***</sup> (0.126)
Never colonized	0.312 (0.160)	0.353 <sup>*</sup> (0.165)	0.229 (0.190)	0.183 (0.174)	0.209 (0.161)
Proportion in largest ethnic group	0.141 (0.200)	0.125 (0.191)	0.135 (0.199)	0.167 (0.192)	0.164 (0.190)
Political freedoms	0.092 <sup>***</sup> (0.033)	0.081 <sup>*</sup> (0.035)	0.078 <sup>*</sup> (0.032)	0.056 (0.033)	0.061 (0.033)
Constant	-4.702 (0.575)	-4.436 (0.565)	-6.097 (0.865)	-5.976 (0.883)	-4.805 (0.564)
Adj. $R^2$	0.75	0.75	0.73	0.76	0.77
SEE	0.48	0.48	0.49	0.46	0.46

Sample size is 93. Mean of dependent variable is 0.18. White-corrected standard errors are shown in parentheses.

<sup>\*</sup> Indicates significance at 0.05 for two-tailed tests

<sup>\*\*\*</sup> Indicates significance at 0.01 for two-tailed tests

the graft index. Each one-point improvement on the political freedoms' index is associated with an increase of nearly one-tenth of a standard deviation on the graft index.

In Eq. (2), the women's influence variable is the share of top ministerial/bureaucratic positions held by women. This variable is also highly significant, and its coefficient is nearly identical to that for women in parliament in Eq. (1). Eq. (3) substitutes women's share of the labor force, which is also highly significant, with a coefficient very similar to those for the women's influence variables in Eqs. (1) and (2). A standard deviation increase in women's share of top ministerial/bureaucratic positions, or in women's share of the labor force, is associated with increases of about one-fifth of a standard deviation in the graft index.

Eq. (4) includes all three women's influence variables: labor force is significant, and parliament and ministers are jointly significant. A possible way of interpreting these results, where women's labor force share and women's participation at elite levels of the government are each significant, is that the former captures women's influence in reducing petty corruption, while the latter captures primarily women's impact on reducing grand corruption. Of course, confirmation of this conjecture requires more detailed information.

Because the three women's influence variables are correlated and each can be interpreted as being a partial measure of the larger concept of women's participation in public life, there is a certain logic for constructing an overall index of women's participation from the separate indicators. We created such an index by standardizing the three variables (mean 0, standard deviation equal to 1) and taking the mean. The index is correlated with its parliament, ministers/bureaucrats, and labor force components at 0.87, 0.84, and 0.66, respectively.<sup>26</sup> The index ranges from a low of -1.91 (UAE) to a high of 3.02 (Finland). In Eq. (5), the women's participation index has a *t*-statistic of nearly 6, and the adjusted  $R^2$  indicates a slightly better fit than in Eqs. (1)–(4). A standard deviation increase in the index is associated with an increase in the graft index of three-tenths of a standard deviation. Fig. 1 depicts the partial relationship, corresponding to Eq. (5).<sup>27</sup>

In Section 2, we saw that the percentage of women who say corruption is "never justified" is higher than the percentage of men who give the same response. It can be argued that greater participation by women in public life should have a larger impact on corruption in countries in which the gender gap (the percentage of women who say corruption is never justified minus the same

<sup>26</sup> Alpha, a measure of index reliability, equals 0.70 for this index. Alpha varies from 0 to 1, and increases with the degree of intercorrelation among the index items and with the number of items in the index.

<sup>27</sup> The slope of the least-squares line in Fig. 1 is 0.364, the coefficient for the women's participation index in Eq. (5).

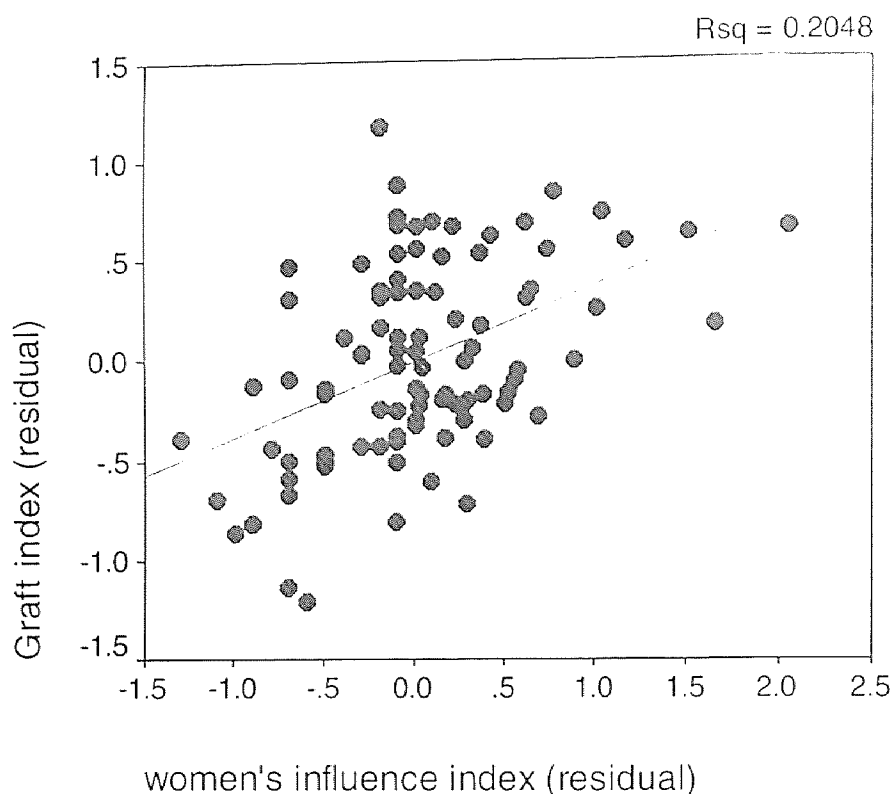


Fig. 1. Graft index and women's influence index (partial plot).

percentage for men) is larger. To test this, we need to estimate a model in which the women's participation variable and the gender gap in attitudes are entered individually and also interacted; the coefficient on the interaction term should be positive. We use data on 32 nations (from the 93 nations in our cross-country sample) for which the World Values Surveys could be used to compute the gender gap in attitudes toward bribery.

The regression reported in column 1 of Table 6 is identical to Eq. (5) of Table 5, with the exception that it adds as regressors the WVS gender gap<sup>28</sup> and its interaction with the women's participation index. The interaction between the gender gap in attitudes and the women's participation index is positive and significant at the 0.07 level (two-tailed test). The women's participation index has a coefficient of 0.24 conditional on a gender gap of 0, but increases by 0.034 for each percentage-point increase in the gender gap. These results are broadly

<sup>28</sup> For countries represented in both WVS survey waves, the gap is the mean of the gaps for 1981 and 1990.

Table 6  
Determinants of corruption, cross-country regressions

Dependent variable	Equation		
	1 Graft	2 Graft (WLS)	3 TI
Women's participation index	0.240 (0.128)	0.380 * * (0.070)	1.187 * * (0.158)
Log (GNP per capita, 1995)	0.733 * * (0.227)	0.606 * * (0.069)	1.660 * * (0.249)
Average years of schooling, 1990	– 0.066 (0.051)	– 0.009 (0.030)	– 0.076 (0.088)
Catholic proportion	0.025 (0.261)	– 0.131 (0.163)	– 0.645 (0.493)
Muslim proportion	– 0.325 (0.295)	0.021 (0.185)	0.426 (0.524)
Former British colony	0.565 * (0.240)	0.480 * * (0.123)	0.840 * (0.391)
Never colonized	0.203 (0.246)	0.144 (0.149)	– 0.372 (0.428)
Proportion in largest ethnic group	– 0.085 (0.824)	0.178 (0.235)	0.806 (0.591)
Political freedoms	0.092 * * (0.033)	0.076 * (0.032)	0.203 (0.110)
WVS gender gap	– 0.016 (0.025)		
WVS gap × participation index	0.034 (0.018)		
Constant	– 6.161 (1.288)	– 5.473 (0.469)	– 10.368 (1.571)
Adj. $R^2$	0.81	0.80	0.81
N	32	93	68

White-corrected standard errors are shown in parentheses. Note that  $R^2$  does not have its usual interpretation in WLS.

\* Indicates significance at 0.05 for two-tailed tests.

\*\* Indicates significance at 0.01 for two-tailed tests.

consistent with the idea that the impact of women's participation on levels of corruption increases as the gender gap in attitudes becomes larger. However, this issue can be better explored when data on gender gaps in the attitude towards bribery become available for more countries.

#### 4.3. Robustness of results

The graft index is constructed from numerous underlying sources. Kaufmann et al. (1999) report standard errors for each country estimate. These standard errors increase with the level of disagreement among sources regarding the severity of corruption in the country, and decrease with the number of sources available for the country. Higher standard errors reflect greater uncertainty about the "true" level of corruption. In Eq. (2) of Table 6, the (inverse of the) standard errors are used to weight observations, to reduce the sensitivity of estimates to the inclusion of countries for which there is less reliable information on the severity of corruption. Results from this weighted least squares regression are very similar to those from OLS in Eq. (5) of Table 5, for the women's participation index and for the other regressors.

The TI index is based on a different approach to data quality issues, requiring that a minimum of three sources be available for a nation, for a value to be

published. The TI index therefore covers fewer countries. Eq. (3) of Table 6 replicates Eq. (5) of Table 5, but substituting the TI index for the graft index. Despite the change in sample size from 93 to 68, results are very similar: the same variables that were significant using the graft index are also significant using the TI index. A standard deviation increase in the women's participation index is associated with an increase in the TI index of two-fifths of a standard deviation, or about one point on the 0–10 TI scale.

Results are also not sensitive to the inclusion of outlying observations. To conserve space, Table 7 reports only the coefficients and standard errors for the women's participation variables, from regressions analogous to Eqs. (5) and (1)–(3), respectively in Table 5, using the graft index. Each row of Table 7 thus summarizes the results from four regressions. The first row duplicates the coefficients and standard errors for the women's participation variables from Eqs. (5) and (1)–(3), respectively. The next two rows of Table 7 show that when we run median or robust regressions, which downweight the influence of outliers, results for the women's participation variables are unaffected.

The fourth set of coefficients in Table 7 is estimated deleting Denmark, Finland, Norway and Sweden, which all rate at or very near the top on all of the women's participation variables as well as on the graft index. The women's variables all remain significant, with magnitudes minimally affected.

Results are also insensitive to the addition of several regressors not included in Table 5. Treisman (1998) found that federal states were more corrupt than nations with centralized governments. We use a dummy variable from Gurr's Polity 98 data set, which classifies 16 of the 93 nations in our sample as federal states. The fifth row of coefficients in Table 7 indicates that the inclusion of this federalism dummy has no impact on the women's participation results. Similarly, adding a trade openness dummy (for 1994, using data from Sachs and Warner, 1996) or the black market currency exchange premium, as measures of opportunities for bribe seeking, does not substantially alter the estimated impact of women's participation on corruption.<sup>29</sup> It is often claimed that public officials are more likely to seek bribes when they are poorly paid (e.g. Haque and Sahay, 1996). Therefore, we include the average government wage as a multiple of per capita GDP.<sup>30</sup> Civil service pay turns out to be unrelated to corruption in our regressions, and its inclusion does not materially affect the women's participation coefficients, despite a reduction in sample size from 93 to 64.

Ethnic homogeneity was not significant in Table 5; an alternative indicator of social cohesion that might be related to corruption is the distribution of income.

<sup>29</sup> Government consumption as a share of GDP also proved to be unrelated to corruption, and its inclusion did not affect the estimates on the women's participation variables.

<sup>30</sup> These data were assembled for the early 1990s by Schiavo-Campo et al. (1997).

Table 7  
Gender and the graft index: alternative samples, specifications, and methods

Change in method, sample, or specification	Index of women's participation	Women's share in parliament	Women's ministerial positions	Women's share in the labor force
From Table 5: Eqs. (5), (1)–(3)	0.364** (0.062)	2.456** (0.751)	2.432** (0.567)	2.419** (0.767)
Median regression	0.335** (0.101)	3.185** (1.107)	3.413** (1.046)	2.965 (1.612)
Robust regression	0.370** (0.086)	2.702** (0.745)	2.444** (0.747)	2.656** (1.022)
Scandinavia dropped ( $N = 89$ )	0.405** (0.092)	2.197* (1.063)	2.559** (0.913)	2.264** (0.773)
Federalism dummy added	0.363** (0.061)	2.424** (0.752)	2.438** (0.561)	2.437** (0.768)
Trade openness dummy added	0.375** (0.064)	2.539** (0.791)	2.508** (0.585)	2.434** (0.759)
Black market exchange premium added ( $N = 83$ )	0.334** (0.061)	2.452** (0.620)	1.861** (0.579)	2.813** (0.660)
Civil service pay added GDP ( $N = 63$ )	0.330** (0.061)	2.516** (0.680)	2.455** (0.591)	1.898** (0.927)
Continent dummies added	0.248** (0.069)	1.731** (0.646)	1.610** (0.579)	1.590 (0.812)
Life expectancy gap added	0.344** (0.063)	2.349** (0.763)	2.504** (0.567)	1.906** (0.793)
Education attainment gap added ( $N = 88$ )	0.317** (0.061)	2.293** (0.680)	1.914** (0.527)	2.359** (0.803)
Herfindahl index of party representation added ( $N = 86$ )	0.358** (0.063)	2.425** (0.725)	2.534** (0.573)	2.222** (0.866)

Cell entries indicate coefficients and standard errors for women's participation variables. Dependent variable is the graft index from Kaufmann et al. (1999). Except where indicated, independent variables are the same as in Table 5. Standard errors are white-corrected, except in median and robust regression.

\* Indicates significance at the 0.05 level for two-tailed tests.

\*\* Indicates significance at the 0.01 level for two-tailed tests.



Including the Gini measure of income inequality has little impact on the women's participation coefficients (not reported in Table 7) and they all remain significant.

We added a set of continent dummies (Sub-Saharan Africa, Middle East and North Africa, Latin America and the Caribbean, and Asia, with the OECD as the omitted category), to account for any omitted variables related to corruption or women's participation rates that vary primarily across continents or country groupings. For example, it is conceivable that the low corruption, high women's participation countries are all developed countries, and that corruption and women's participation are unrelated within the group of developed countries, or within the group of developing countries. As shown in Table 7, the inclusion of the continent dummies reduces the women's participation coefficients, but all remain significant (labor force only at the 0.06 level).

It could be that more corrupt countries also discriminate more against women, which leads to lower levels of participation by them. In this scenario, the observed correlation between women's participation and corruption is spurious, and driven by the omitted variable "level of discrimination against women". We evaluated (and ruled out) this possibility by controlling for the level of gender discrimination using the gap between men's and women's educational attainment, and the gap between men's and women's life expectancy. Inclusion of these controls changes the women's participation estimates only slightly, as shown in Table 7. In results not shown in Table 7, results for the women's participation variables also proved insensitive to the inclusion of Humana's (1992) indexes of "political and legal equality" and "social and economic equality" between men and women.

Studies of the determinants of the presence of women in parliament find that proportional representation matters (Reynolds, 1999; Matland and Studlar, 1996). In a PR system as opposed to a plurality system, a certain number of women can be elected even if a large majority of voters in every district is disinclined to vote for women candidates. Electoral rules can also potentially affect the incidence of corruption. Myerson (1999) argues that it is easier under PR than under plurality voting to launch a new party that adopts policy positions similar to those of an existing party, but which promises to reduce corruption. On the other hand, as Myerson notes, reforms in Italy to reduce the number of seats allocated through PR are motivated by the goal of reducing corruption, apparently because PR is blamed for the Christian Democrats' long-standing dominance of the governing coalition. Thus, because proportional representation can affect the incidence of corruption as well as the presence of women in parliament, its omission could bias our estimates. Using data from Beck et al. (2000), we identify 38 nations in our 93-country sample that elect the majority of their (lower house) legislators via PR. A PR dummy when added to our regressions is associated with a one-fifth of a standard deviation decline in the graft index (significant at 0.10 for a two-tailed test), but this does not reduce the magnitude or significance of the women in parliament coefficient (2.878, S.E. of 0.680).

Reynolds (1999) also finds that more fragmented multi-party systems are associated with lower women's representation in parliament, attributing this result to the likelihood that smaller parties have a smaller pool of "safe seats" for which they can nominate women candidates. If the extent of fragmentation affects the level of corruption as well as women's representation, its omission is a potential source of bias. A Herfindahl index of party representation in the legislature (from Beck et al., 2000) is not significant when added to our corruption regressions, and its inclusion does not affect the relationship between corruption and women's share of parliamentary seats, as can be seen in the last row of Table 7.

Four countries in our sample have minimum quotas or reserved seats for women in parliament, and 22 more had some political parties with quotas. Dummy variables for these national and party quotas are not significant in our corruption regressions, and their inclusion has no impact on the women in parliament estimates.<sup>31</sup>

Despite our efforts, the possibility of omitted variable bias in a cross-sectional regression can never be entirely ruled out. For example, it might be argued that societies in which "traditional" political values and clientelistic attitudes are prevalent may tend to choose "strong" male leaders, and may also be more tolerant of corruption, with no causal connection between these two phenomena. A powerful way of addressing this problem of potential omitted variable bias is by looking at the correlation between the changes in women's participation and the extent of corruption *within* countries over time; we can then be confident that our estimates are not affected by the omission of any time-invariant country-specific variable that affects the level of corruption. Time-series data on corruption are very sparse: the TI index is updated annually but was first released only in 1995. The graft index is even newer, and has been produced for only 1 year thus far. The only available source covering a substantial period of time and a large sample of countries is the "corruption in government" index from the International Country Risk Guide (ICRG). The ICRG corruption index, one of the sources used in constructing the TI index and the graft index, varies from 0 to 6, with higher

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<sup>31</sup> Countries with quotas are identified from Inter-Parliamentary Union (1997), which conducted a survey of national parliaments in 1996. However, some parliaments did not respond to the survey. In most cases, the dates of quota adoption are not available, and some of these quotas may not have been in effect when parliaments of 1994, which is the most recent year for which we have data, were elected. Htun and Jones (1999) discuss why quotas have had limited effectiveness in increasing women's representation. Other determinants of women's representation in parliament have been identified in the political science literature (see Reynolds, 1999 and Darcy et al., 1994). However, these determinants, such as the date women were first granted the vote or the right to run for office, can be viewed as alternative and less precise measures of women's participation in the context of the current study, rather than as omitted variables that could influence the relationship between women's participation and corruption.

Table 8

Determinants of corruption changes, cross-country regressions (Dependent variable: change in ICRG corruption index from 1982 to 1997)

Equation	1	2
Parliament, percentage women (1994 minus 1975)	3.948 * * (1.531)	
Labor force, percentage women (change in)		7.603 * * (3.458)
Percentage of change in GNP per capita, 1982–1996	0.124 (0.295)	0.086 (0.275)
Change in political freedoms	0.084 (0.056)	0.086 * (0.042)
Initial value of corruption index	–0.607 * * (0.057)	–0.577 * * (0.057)
Constant	1.956 (0.208)	1.742 (0.199)
Adj. $R^2$	0.51	0.51
$N$	85	98

\* Indicates significance at the 0.05 level for two-tailed tests.

\*\* Indicates significance at the 0.01 level for two-tailed tests.

values representing less corruption.<sup>32</sup> Correlations of the ICRG corruption index for 1997 with the TI and graft indexes, respectively, are 0.80 and 0.77. Because the TI and graft indexes aggregate information from numerous sources, the ICRG index, based on a single source, is likely to be a noisier measure.

Time-series data are available on two of the women's participation variables. Women in the labor force is available annually. Women in parliament is available for 1994, 1990, 1985 and 1975. Time-series data on our control variables are available for two of the three control variables that are statistically significant in Table 5: per capita income, and the political freedoms index. The other one—the ex-British colony dummy, does not vary over time. Most of the other control variables in Table 5 (e.g. percentage of Catholics) are unlikely to vary much over a 15-year period.

The dependent variable in our time-series tests in Table 8 is the change in the ICRG corruption index from 1982 (or the first available year) to 1997. Changes vary from –4 (Niger) to +3 (Bahamas, Haiti, Iran, Philippines and Syria). Control variables include the initial corruption index value, to control for regression-to-the-mean effects, the percentage change in per capita income over the period, and the change in the political freedoms measure. Women's participation variables are the change in women's share of parliamentary seats (Eq. (1)) and the change in women's share of the labor force (Eq. (2)). The former varies from –27 percentage points (Albania) to +25 points (Guyana), while the latter varies from –2.3 percentage points (Botswana) to +12.1 points (Kuwait).

The ICRG index shows a strong regression-to-the-mean effect, as the initial value of the index has a coefficient of –0.6 and a  $t$ -statistic of about –10 in both

<sup>32</sup> We follow others who have used this measure in using the annual values calculated by Knack and Keefer (1995) from the monthly ICRG issues dating back to 1982.

regressions. Increases in income are not associated with significant improvements in corruption ratings. Change in political freedoms has a positive coefficient as expected, but it is significant only in Eq. (2). Coefficients for both women's participation variables are positive and significant. In Eq. (1), an increase of 25 percentage points in women's share in parliament is associated with a one-point improvement in the ICRG corruption rating. In Eq. (2), an increase of about 13 percentage points in women's share of the labor force is associated with a one-point improvement in the ICRG corruption rating. These results suggest that our central finding is not driven by a time-invariant, country-specific omitted variable. Of course, we still cannot entirely rule out bias due to a time-varying omitted variable (say, social "modernization").

## 5. Interpretation of findings and concluding remarks

Though questions can be raised regarding each of the three sets of evidence we have assembled, they reinforce each other, and taken together, make a strong case. For instance, the World Values Survey results can be criticized on the grounds that they represent hypothetical choices, and the data on corruption from Georgia can be questioned because they are self-reported. However, neither of these charges holds true for the cross-country data. The results from Georgia can be questioned on grounds of selection bias; if employers discriminate against women, only those women who are exceptionally capable or honest may become owners/managers, and the gender differential we are observing may be the difference between average men and exceptional women. Arguing along similar lines, greater representation of women in government and in market work could improve average outcomes because participation rates for women are still low, and women participants are from the "better" part of the women's distribution, rather than because the distribution of attitudes towards corruption differs between men and women. However, in the World Values Survey, we have random samples of the whole population (no room for selection bias) and there is still a gender differential. Moreover, controlling for discrimination against women does not change our cross-country results. We are making a simple point: to question the central finding of this paper, one needs to argue that the results of careful analyses of several distinct data sets have, by sheer fluke, all been biased in the same direction. Our conclusion, that there is indeed a gender differential in tolerance for corruption, is more plausible.

We are reassured to learn that our evidence is entirely consistent with the findings of leading criminologists. For instance, Gottfredson and Hirshi (1990, p. 194) show, using U.S. Department of Justice figures, that arrests for embezzlement per 100,000 white-collar workers are higher for men for every age group. They also cite a variety of sources to make the case that across age groups, countries, and types of crime, the evidence regarding higher participation of men is remark-

ably uniform. The following summary statement from a study conducted by the National Academy of Sciences of the United States reflects the confidence with which the gender differential has been identified in the criminology literature:<sup>33</sup> “The most consistent pattern with respect to gender is the extent to which male criminal participation in serious crimes at any age greatly exceeds that of females, regardless of the source of data, crime type, level of involvement, or measure of participation.”

This paper has primarily focused on identifying an empirical regularity, a “stylized fact”. Ideally, we would like to conclude with an explanation of this stylized fact. Criminologists have developed many theories that are potentially relevant. Women may be brought up to be more honest or more risk averse than men, or even feel there is a greater probability of being caught (Paternoster and Simpson, 1996). Women, who are typically more involved in raising children, may find they have to practice honesty in order to teach their children the appropriate values. Women may feel more than men—the physically stronger sex, that laws exist to protect them and therefore be more willing to follow rules. More generally, girls may be brought up to have higher levels of self-control than boys which affects their propensity to indulge in criminal behavior (Gottfredson and Hirshi, 1990, p. 149).

Though these theories are generally consistent with our findings, our data are not sufficiently detailed to specifically support any one of them. Indeed, even in heavily researched areas such as male–female differences in sexual behavior and propensity for violence, researchers are very far from having reached a consensus on the underlying causes. Since at least the mid-1970s, sociobiologists have argued that behavioral differences between human males and females have parallels among other species, and that common explanations can be provided for these differences, based primarily on the roles of males and females in reproduction and child-rearing.<sup>34</sup> However, many sociologists, especially those of feminist persuasion, have seen in these arguments the potential for biological justification of gender inequity; some have accused sociobiologists of “biological essentialism”, i.e. an emphasis on biology to the point where culture is treated as peripheral to behavior, and the social construction of gender roles is severely underestimated (Bem, 1993; Epstein, 1988). However, there is disagreement even among feminists; some radical feminists have implicitly appealed to biology to theorize a greater propensity for nurturing and co-operative behavior in women, following from their role in reproduction and child-rearing.<sup>35</sup> Despite decades of debate, these issues are far from resolved; it is clear that empirical identification of the

<sup>33</sup> Blumstein et al. (1986), cited in Gottfredson and Hirshi (1990, p. 145).

<sup>34</sup> Early influential works in sociobiology include Wilson (1975, 1978). Pinker (1997) provides an accessible summary.

<sup>35</sup> Jaggar (1983) provides a critical overview of various feminist perspectives.

sources of gender differentials in behavior is a very difficult (not to mention politically charged) task even for researchers who have much richer sources of data than we do. Therefore, though there is a plethora of theories regarding the sources of gender differentials in crime, with potential applications to corruption, we are reluctant to take a position on this issue.

We do need to comment, however, on one policy-related matter. If gender differentials in tolerance for corruption are culturally based, it is worth asking whether they will persist as the position of women in society changes and their participation in the labor force increases. We suspect the differentials will persist, at least in the medium term, for three reasons. First, in our evidence from the World Values Survey, the gender differential is robust to controlling for employment status. Second, in the same survey, the gender differential can also be seen in the OECD countries, where women have been in the labor force in large numbers for some decades. Finally, it used to be routinely assumed by criminologists that with greater equality of status between men and women, crime rates would equalize. However, in the United States, large differentials have persisted despite the increase in women's participation in the labor force (Gottfredson and Hirshi, 1990, pp. 146–147). Given this evidence, we suspect the gender differential in corruption will be stable in the medium term, and policy initiatives like those discussed at the beginning of this paper will indeed reduce corruption.

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